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Abstract: Federal Supplemental Compensation (FSC) was the program that temporarily extended the duration of Unemployment Insurance (UI) benefits by 8 to 14 weeks during 1982-1984. This paper examines whether and to what degree the extension of benefits under FSC increased the expected length of UI recipients' jobless spells. The estimates are derived from a large UI-administrative data base that spans the expiration of FSC and allows one to observe whether and (approximately) when a worker actually returned to work. Previous studies of the effects of extended benefits have had to make assumptions about UI recipients' return to work that appear to yield downward-biased estimates of the effect of extended benefits on expected jobless duration. The estimates presented here suggest that an additional week of benefit eligibility increases a UI recipient's expected spell of joblessness by nearly one week. Moreover, the estimates suggest that claimants who exhausted their regular UI benefits and were ineligible for an additional 12 weeks of FSC were more than six times as likely to return to work as claimants who exhausted their regular benefits but were eligible for FSC. Hence, the findings offer striking evidence that workers tend to find jobs just as their UI benefits expire.

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Stephen A. Woodbury

Since the mid-1970s, the most-researched question about the Unemployment Insurance (UI) system has been whether and to what degree higher weekly UI-benefit amounts lengthen UI recipients' jobless spells. But an equally important and under-researched question is how the potential duration of those benefits influences the length of jobless spells. The latter question is important for two reasons. First, since the 1950s, a variety of extended UI-benefit programs have been legislated at the federal level, making the potential duration of UI benefits a highly variable aspect of the UI system. Second, as will become clear, econometric problems make far more tenuous any inference about the influence of UI-benefit extensions on the expected length of UI recipients' jobless spells.

This paper explores the effects of the most recent UI-benefit extension program, the Federal Supplemental Compensation (FSC) program, which temporarily extended the duration of Unemployment Insurance (UI) benefits by 8 to 14 weeks during 1982-1984. Estimates of how FSC influenced the length of UI recipients' expected jobless spells are derived from a large UI-administrative data base that was gathered just before and just after the expiration of FSC in late 1984. These data offer two advantages over the administrative data used in most existing studies of the effects of potential benefit duration on expected jobless duration. First, because the data were gathered just before and after the expiration of the FSC program, eligibility for FSC depends (for most of the sample) only on the timing of a worker's initial claim for UI benefits, and is independent of each worker's characteristics and earnings history. Second, the data allow one to observe whether and (approximately) when a worker actually returned to work. The ability to observe actual return-to-work allows one to overcome what appears to be the most important econometric problem involved in appraising the effects of extended benefits on expected jobless duration--that of correctly specifying whether an observed spell of joblessness is complete or censored.

The next section offers some background--both historical and economic--on extended benefits. Section II characterizes the problems of measurement that appear to have hampered researchers in their efforts to obtain unbiased estimates of the influence of potential benefit duration on the expected length of UI recipients' jobless spells. That section also treats the data used in this study and describes how they can yield improved estimates of the effects of extended benefits. Section III develops and presents estimates of several models that allow one to infer the effects of extended benefits. Included are a simple linear duration model, a parametric jobless duration model that accounts for censoring of the dependent variable and non-normality of the error term, and a semi-parametric model of the conditional probability (or hazard) of returning to work. A final section summarizes the findings of the paper.

I. Potential UI Benefit Duration and Extended Benefits

A. Institutional and Policy Background

At the outset of the U.S. Unemployment Insurance program in 1936, the decision was made to limit benefit duration, mainly on the grounds that benefits that could not be financed out of anticipated contributions should not be provided (Haber and Murray 1966, pp. 111-113). That decision, which today is so widely accepted that it is taken for granted, has led to a variety of policy questions and economic issues. For example, on what grounds is benefit duration to be limited, and what should the limit be? Should potential duration of regular benefits be uniform for all workers, or vary with a worker's employment history? Should extended benefits be made available to workers who exhaust their regular benefits during economic downturns? If so, how should the duration of the extended benefits be set? Should extended benefits be discretionary--that is, made available when Congress and the Administration see a need--or triggered automatically by prespecified macroeconomic conditions? If an automatic trigger is chosen, how should the prespecified conditions be set?

Today, limited ~~limited~~ benefit duration is accepted as a premise of the UI system both for financial reasons (as in 1936) and because it is widely believed that to make available benefits of unlimited duration would weaken incentives to seek employment.

This latter point--that longer potential benefit duration leads to longer spells of joblessness--is now the central policy issue in regard to the duration of benefits for the obvious reason that answers to the six questions posed above depend on it.

Even though limited benefit duration is accepted as a premise of the UI system, it has become clear in the last 20 years that the UI system is expected to respond to severe cyclical downturns by extending benefits for workers who have exhausted their regular benefits. Indeed, Murray (1974) has stated this as a premise of the UI system, and both Hight (1975) and Corson and Nicholson (1982) have shown how extending benefits during a recession can alleviate hardship and reduce exhaustion rates to levels as low as would occur under good economic conditions.

The initial rationale for extended benefits was that as labor market conditions worsen during a recession, jobless spells lengthen as it becomes more difficult to find a job. This basic rationale was undermined during the 1970s by research showing that most jobless spells are of short duration, even those spells experienced during a recession (Perry 1972; Hall 1972). Indeed, recent evidence reported by Dynarski and Sheffrin (1987) suggests that the average duration of a jobless spell varies little over the business cycle. Such findings call into question the soundness of extending benefits during cyclical downturns.

Three types of findings have mitigated the effects of the arguments against extending benefits during downturns. First, various studies of workers who have exhausted their UI benefits (Mathematical Policy Research 1976; see Murray 1974 and Hamermesh

1977 for reviews of the earlier exhaustee studies) suggest that exhaustees are strongly attached to the labor force, and are not workers who would have dropped out of the labor force had they been ineligible for benefits. Second, the turnover view of the labor market--that most jobless spells are short--seems to have neglected the existence of a relatively small group of workers who experience very long unemployment spells (Akerlof and Main 1980; Clark and Summers 1979). Third, and most important, the most recent studies of workers' responses to extended benefits suggest that extending the potential duration of UI benefits does not appreciably increase the expected duration of jobless spells (see in particular Moffitt 1985a, 1985b). If these latter findings are accepted, then there can be little objection to extending UI benefits during cyclical downturns.

B. Theoretical Issues

Most estimates of the effects of both benefit duration and benefit amounts on the duration of joblessness have been based on one or another model of job search (see Mortensen 1986). These models provide a theoretical link between the duration of joblessness, on the one hand, and job-search intensity, individual characteristics, and labor market conditions, on the other. It is useful to review a general job-search model as a prelude to the empirical work.[1]

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1. Other models also provide possible bases for empirical work on unemployment duration. Moffitt and Nicholson (1982) have used the familiar income-leisure model of labor supply, and Davidson

Let T denote the week in which a UI recipient returns to work, and let P_t denote the probability that a UI recipient returns to work in week t, given that she has not already returned to work by then; that is, $P_t = Pr[T=t|T \geq t]$. Then we can express P_t as the product of (a) the probability of receiving a job offer in week t and (b) the probability of accepting that job offer, given that an offer has been made. The probability of receiving an job offer in week t (J_t) depends on the intensity of the worker's job search (i) and a vector of characteristics of the worker that determine the demand for the worker's labor (c); that is, $J_t = J_t(i,c)$. The probability of offer acceptance (A_t) depends on whether the offered wage (w^o) equals or exceeds the worker's reservation wage (w^r); that is, $A_t = Pr[w^o \geq w^r | J_t]$. Hence, we can express the probability of finding reemployment during week t (given that reemployment has not already occurred) as:

(1) $P_t = J_t(i,c)\ A_t.$

We have defined P_t in discrete time above. In the limit, as the time interval over which reemployment is measured approaches zero, P_t becomes an instantaneous rate of reemployment known as a hazard rate, h(t). The hazard rate is linked to unemployment duration in the following way. If t has cumulative distribution F(t), and frequency distribution f(t), then $h(t) = f(t)/[1 - F(t)] = f(t)/S(t)$, where S(t) is the so-called survivor function, or the probability of being unemployed to time t (Lancaster 1979). We can also express the survivor function as dependent on the hazard:

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and Woodbury (1989) have used a job-matching model.

$S(t) = f(t)/h(t)$. Thus, the hazard rate is inversely related to the survival probability, and any factor that increases the hazard rate should decrease expected unemployment duration. For example, equation (1) highlights the fact that longer spells of joblessness can result from less-intense job search, from individual characteristics (c) that imply lower demand for a worker's services, or from a lower probability of job-offer acceptance. From the point of view of the present work, the probability of offer acceptance, A_t, and the intensity of job search, i, are central, depending as they do on the generosity and duration of UI benefits.

II. Problems of Data and Inference

The data used here come from the administrative records of the State of Illinois Unemployment Insurance system, and were gathered as part of the Illinois UI Bonus experiment (see Woodbury and Spiegelman 1987, pp. 514-518, for a fuller description). The workers examined here are the 2,162 UI-eligible male claimants who were assigned to the control group of the experiment.

The Illinois data are typical of the data used in past studies of the effects of extended benefits on the duration of unemployment because they are administrative data--that is, data gathered and maintained by agencies responsible for administering the UI program. Indeed, it would make little sense to use any other type of data to address this issue, because only in administrative data is there accurate information on a claimant's eligibility for UI and extended benefits.

But the Illinois data have three features that make them unique among UI administrative data, and especially well suited to an examination of the influence of potential benefit duration on expected jobless duration. First, they span the expiration of the FSC program at the end of 1984. Consequently, they permit one to compare the jobless spells of workers who were actually eligible to FSC with the spells of workers who would have been eligible if they had become unemployed and filed for benefits only weeks earlier. In other available data sets, eligibility for extended benefits often depends on an individual's work history (base period earnings or weeks of employment prior to layoff). Inferences about the independent contribution of potential benefit duration to the length of a jobless spell should be less tenuous with the Illinois data than in data where potential duration is correlated with work history.

Second, the Illinois data include both earnings history data and a return-to-work indicator, which permit one to make an accurate classification of jobless spells as complete or censored.[2] Earlier researchers have frequently had to make what appear to be erroneous assumptions about whether a spell of insured unemployment represents a complete or censored spell of

2. Because these are administrative data, there is no way of distinguishing unemployment from out-of-labor-force status for workers who have no earnings after the spell of insured unemployment ends. Accordingly, we focus on the duration of joblessness (meaning either unemployment or out-of-labor force status), and with the probability of return to work.

joblessness. A more complete discussion of this issue and its consequences follows.

A. Censoring Problems

Most analyses of the effects on jobless duration of potential benefit duration and benefit amounts have used basic benefits data from UI administrative files. These data are extremely rich: For example, in the Illinois data, the so-called Benefits Information System contains demographic data on claimants, the dates of their UI claims, and the amount and timing of benefits received.

But basic benefits data from UI administrative files are usually deficient in that they exclude any information on the subsequent earnings of claimants. Hence, they fail to offer data on actual spells of unemployment. Rather, they indicate only the duration of <u>insured</u> unemployment experienced by a claimant.

In the Illinois data, this deficiency can be partly overcome by two additional pieces of data. The first is a return-to-work indicator that was constructed with the cooperation of Illinois Department of Employment Security personnel during and after the experiment that generated the data. The second is a separate administrative data base known as Wage Records, which stores information on the earnings history of workers both before and after their spell of insured unemployment. By matching each claimant's wage record to his or her benefits data, it is possible to determine whether a spell of insured unemployment was followed by a period of earnings. If the observed spell of insured unemployment was followed by a period of earnings, and if the return-to-work indicator was positive, then we can infer that the insured spell and the actual spell of joblessness were the same. On the other hand, if the insured spell was not followed by a spell of earnings, the insured spell must be considered a censored or truncated measure of the actual spell of joblessness.

It is rare for administrative data to have a return-to-work indicator or Wage Records matched to the benefits data, as in the Illinois data. As a result, existing research using administrative data has necessasrily taken a different approach to drawing inferences about actual spells of joblessness from observations on insured unemployment spells. First, because the number of weeks of unemployment that can be observed in administrative data is limited by the potential duration of UI benefits, it has been assumed that any spell of insured unemployment that is at the maximum potential is an incomplete spell. For example, a worker eligible for 26 weeks of state regular benefits who is observed receiving 26 weeks of UI benefits would be considered to have had a jobless spell of greater than 26 weeks. Second, and conversely, a workers eligible for 26 weeks of benefits who is observed receiving less than 26 weeks of UI benefits would be considered to have had a jobless spell of exactly the observed length.

Neither of these assumptions is necessarily correct, as can be seen in Table 1, which shows four cases, labeled A through D. Cases A and B are those of workers who received the maximum potential weeks of UI benefits--that is, exhausted their benefits.

It is possible for such workers to return to work immediately after receiving their last benefit payment (Case A), or to continue to be out of covered employment (Case B, which implies either continuing to seek employment or dropping out of the labor force after receiving the last benefit payment).[3] The usual assumption is that all workers who exhaust benefits continue without covered employment, as in Case B. But the right-most column of Table 1 shows that this assumption is incorrect for nearly 40 percent of the men in the control group of the Illinois UI experiment who exhausted their benefits. That is, the data show that 283 of the 717 men who exhausted their benefits returned to work immediately (or very shortly) after receiving their last benefit payment.

Likewise, it is possible to misclassify a worker who did not exhaust his or her benefits. Cases C and D in Table 1 are those of workers who received fewer than the potential weeks of benefit payments. Again, the usual assumption is that all such workers returned to work immediately after they stopped receiving benefits, as in Case C. But the right-most column shows that 405 (or 28 percent) of the 1,445 men who ended their benefits before exhausting did not return to covered employment.[4]

3. Obtaining uncovered (usually underground) employment and moving out of state are additional possibilities.

4. Most likely, these men either dropped out of the labor force or took uncovered employment, although it is possible that they stopped participating in UI and continued to seek employment. There is no way of distinguishing these possibilities in the administrative data.

B. Censoring and Inference

Estimating the effects of extended UI benefits on expected unemployment duration offers a fine example of how censored data can interfere with statistical inference. Because the Federal Supplemental Compensation program expired about half-way into the enrollment period of the Illinois experiment, men in the control group of the Illinois experiment can be divided into four categories: (a) those who were eligible for FSC because they met the monetary eligibility criteria for FSC[5] and filed their initial UI claim while FSC was still in effect; (b) those who were monetarily eligible for FSC, but claimed benefits too late to actually receive FSC; (c) those who were monetarily ineligible for benefits, but filed their initial claim before FSC expired; and (d) those who were monetarily ineligible for benefits, and filed their initial claim after FSC expired.

Table 2 shows the mean insured unemployment duration for each of these four groups. What we appear to have is a situation in which the expiration of FSC gives us a natural experiment. The mean unemployment duration of workers eligible for FSC (21.4 weeks) can be compared with the mean unemployment duration of workers monetarily eligible but temporally ineligible because they filed after FSC expired (17.9 weeks). As a quasi-control, the mean unemployment duration of workers who were monetarily

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5. The monetary eligibility criteria for FSC were somewhat more stringent than for state regular benefits. That is, not all UI claimants who were monetarily eligible for state regular benefits were also monetarily eligible for FSC.

<u>in</u>eligible but who filed for benefits while FSC was still in effect (16.5 weeks) can be compared with the mean unemployment duration of workers who were neither monetarily nor temporally eligible (20.1 weeks).

The two comparisons are shown in the bottom row of Table 2 (labeled "Difference"). The difference between the two groups of monetarily eligible workers, 3.5 weeks, suggests that FSC prolonged unemployment spells significantly. Moreover, the difference between the two groups of monetarily ineligible workers, -3.5 (with a large standard error), suggests that there was no underlying macroeconomic or other reason for expecting unemployment spells to be longer after the expiration of FSC. The conclusion would seem to be that workers eligible for FSC tended to take over three weeks longer to return to work than did workers who were not eligible for FSC.

Such an inference would clearly be wrong, though, because it is impossible to observe more than 26 weeks of unemployment among workers ineligible for FSC, whereas we can observe up to 38 weeks of unemployment among workers eligible for FSC. The truncation or censoring of unemployment spells at the maximum potential duration leads to a situation in which the two group means cannot be compared. To take Moffitt's (1985a) extreme example, every worker in each of the two groups might have an actual spell of joblessness of 30 weeks, but we would observe an average of 26 weeks for the first group (because the data are censored at 26 weeks) and 30 weeks for the second (because censoring occurs only at 38 weeks). It may still be, of course, that FSC tended to lengthen jobless spells, but the data in Table 2 cannot be used to make such an inference.

In the presence of censoring, quasi-experimental comparisons like those presented in Table 2 fail to yield reliable estimates of the effects of extended benefits on jobless duration. Hence, other methods of inference must be sought. It is convenient that a variety of methods have been developed for handling censored duration data--methods that allow one to make inferences about actual spells of joblessness from spells of insured unemployment (see the recent review by Kiefer 1988). Some of these methods are applied in parts B and C of the next section.

III. Models of Unemployment Duration and Reemployment Hazard

Estimates of how potential benefit duration affects the expected duration of joblessness have progressed through three stages. In this section, the approach represented by each of these stages is outlined, and estimated results of models based on each approach are presented.

A. A Linear Model of Insured Unemployment Duration

The earliest empirical work on the effects of potential duration on expected jobless duration took the straightforward approach of regressing the duration in weeks of insured unemployment (D) on appropriate explanatory variables (x_1, ...,

x_K), including measures of the replacement ratio and potential duration of benefits:

(2) $D = a_0 + a_1x_1 + \ldots + a_Kx_K + u,$

where u is assumed to be a normally distributed disturbance term. The coefficients of x_1 through x_K provide an estimate of the relationship between the explanatory variables on weeks of insured unemployment. Studies that took this approach include Ehrenberg and Oaxaca (1976), and Holen (1976), among others.

Column (2) of Table 3 displays the results of such a model applied to the Illinois control-group men. The results are in keeping with the simple comparison of means shown in Table 2, in that the coefficient of the FSC eligibility variable suggests that the availability of FSC increased the expected length of jobless spells by over three weeks.

Note that a variety of control variables, which are of secondary importance for present purposes, have also been included in the model: age, ethnicity, the number of employers that the claimant worked for in the base period, base period earnings, the number of referrals received by the claimant from the Job Service, whether a dependents' allowance was received, the weekly UI benefit amount, and the conditions of the labor market in which the claimant was searching for work.

B. Parametric Models of Time to Reemployment

It is now well-known, however, that the assumption that u in equation (2) is normal is untenable, and that application of Ordinary Least Squares (OLS) to equation (2) will produce biased coefficient estimates. There are two reasons for this. First, as already discussed, D is a censored measure of actual jobless duration, since each worker is eligible for a specified maximum number of weeks of benefits. As a result, the distribution of D is truncated at the maximum benefit duration. Realization of this problem lead to some studies that assumed that the underlying distribution of jobless spells is normal, and as a result assumed that the distribution of u in equation (2) is truncated normal. For example, Newton and Rosen (1979) and Classen (1979) both used Tobit analysis--which assumes that u has the truncated normal distribution--to correct for the truncation of the dependent variable. Newton and Rosen's results suggest that an additional week of potential benefit duration leads to 0.4 to 0.5 additional weeks of insured unemployment, whereas Classen's estimates suggest at most an additional 0.1 weeks.

The second reason for questioning the assumption that u in equation (2) is normal is that the empirical frequency distribution of weeks of insured unemployment in most data is not bell-shaped, as the normality assumption requires. Rather, it shows one spike at zero weeks of unemployment, and falling frequencies for greater unemployment durations, until a spike appears where censoring occurs (that is, at maximum benefit duration.) Except for the spike at the censoring point, the empirical distribution looks much like an inverse exponential. This latter problem can be solved in a jobless-duration equation like (2) by making a more appropriate assumption about the distribution of u, and estimating equation (2) under that alternative distributional assumption. The Weibull distribution has been widely assumed in studies of jobless duration because it provides an approximation to the empirical distribution of jobless duration that appears to be valid (Lancaster 1979). (The exponential distribution is a special case of the Weibull. Whereas the exponential restricts the conditional probability that a UI recipient will become reemployed (P_t, or the hazard rate) to be constant over the spell of unemployment, the Weibull allows for the possibility that a UI recipient's probability of reemployment rises or falls over the spell. The greater generality of the Weibull distribution makes it the preferred choice.)

Several studies have imposed a more appropriate distributional assumption on u in equation (2), although few of these have examined the effects of potential benefit duration on the length of unemployment spells. Among the few that do, Solon (1985) has estimated that an additional week of potential benefit duration leads to 0.3 additional weeks of insured unemployment.

Columns (3) and (4) of Table 3 display the results of estimating equation (2) under the assumption that the disturbance term u has the Weibull distribution. The interpretation of the Weibull model's coefficients is straightforward: Each coefficient gives the approximate proportional change in unemployment duration that is attributable to a unit change in the explanatory variable.[6]

The two Weibull models differ in the way that censored observations are defined. In the model displayed in column (3), the usual censoring assumptions are made--all observations that are at the maximum potential benefit duration are defined as censored, and all other observations are taken to reflect actual weeks of joblessness. As discussed in section II.B, this is a misspecification of censoring, and column (3) is labeled accordingly. In the model displayed in column (4), I took advantage of the additional information on subsequent employment available in the Illinois data. This additional information allowed me to correct the misclassification of some completed spells of unemployment as censored (and vice versa), and column (4) is accordingly labeled "Corrected Censoring."

The main difference between the estimates resulting from the two Weibull models is in the estimated effect of FSC on the expected duration of joblessness. This should not be surprising, because the effect of FSC might well be concentrated on workers who have long spells of joblessness. These are precisely the workers about whom we are making crucial assumptions regarding whether a spell of joblessness is complete or censored. From the model of column (3), we can infer that the availability of FSC tends to <u>reduce</u> the expected duration of joblessness. This

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6. The exact proportional change in duration attributable to a unit change in an explanatory variable is $[\exp(b_k) - 1]/\text{shape}$, where shape is the Weibull shape parameter. If b_k is small, and if the shape parameter is close to unity, then this expression approximately equals b_k.

inference should not be taken seriously because there is good reason to believe that in this model, some completed spells have been treated as censored and vice versa.

Censoring has been more accurately specified in the Weibull model of column (4), and greater credence can probably be given to the finding that FSC tends to increase the expected duration of joblessness by nearly 5 weeks.[7] This positive relationship between FSC and jobless duration is also more plausible theoretically, and in keeping with earlier findings.

Hence, the evidence to this point suggest that accounting for censoring of the dependent variable and non-normality of the error term yields estimates of the effects of FSC that exceed estimates derived from simpler techniques that ignore these problems. Also, the evidence suggests that the approach to censoring taken in earlier research tends to yield lower estimates of the effects of extended benefits than does the corrected method permitted by the Illinois data.

C. Semi-Parametric Hazard Models

Two problems cannot be handled in either of the duration models discussed to this point. The first is that some variables may change during a worker's spell of joblessness. For example,

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7. The proportional effect of FSC is found by exponentiating the coefficient of FSC (0.18), subtracting 1, and dividing by the Weibull shape parameter (0.813). This yields a 24 percent increase in duration attributable to FSC. Multiplying 0.24 by the sample mean weeks of observed unemployment (19.75) yields a 4.8 week increase in duration attributable to FSC.

the number of weeks since the initial claim for UI benefits can be thought of as a variable that increases weekly. There is no way of understanding the effects of such "time-varying" explanatory variables in a duration model.

The second is that the duration models force us to make an assumption about the distribution of the error term u in equation (2), and incorrect distributional assumptions will yield misleading inferences about the effects of FSC. For example, the Weibull seems a good approximation to the empirical distribution of jobless spells as long as it is true that the spike in the empirical distribution in the week following benefit exhaustion results from censored data. But if the distribution of jobless spells shows a true spike in the week following benefit exhaustion--that is, if workers tend to put off finding taking a job until just after their benfits terminate--then the Weibull is a poor choice. Ideally, one would like to impose no distributional assumption at all.

To analyze the effects of time-varying explanatory variables and to avoid any assumptions about the distribution of jobless spells requires reconceptualizing the duration problem as a problem of rate of escape from joblessness. In other words, rather than regress some measure of duration on various explanatory variables, one could regress a dummy variable (R_t) equal to one if a worker escaped from unemployment in week t (zero otherwise) on various explanatory variables, some of which are time-invariant (x_1, ..., x_K), and others which are time-varying ($z_1(t)$, ..., $z_N(t)$). Since the dependent variable is a measure of a worker's probability of escaping joblessness in week t (given that the worker was "at risk" of escaping joblessness at the beginning of the week), it is appropriate to interpret the dependent variable as a hazard rate (as discussed above), and to call it h(t). Hence, this model can be written:

$$(3) \qquad R_t = h(t) = b_0 + b_1x_1 + \ldots + b_Kx_K + c_1z_1(t) + \ldots + c_Nz_N(t) + e.$$

If e is assumed normally distributed, then we have a linear probability model, which is used in the exploratory work reported presently.[8]

Each coefficient in equation (3) represents the change in the probability of reemployment that results from a unit change in the independent variable. For example, if x_1 were age in years, then b_1 would show the change in probability of reemployment associated with an additional year of age. This change in probability would be assumed constant over the spell of unemployment (unless age were interacted with a time-varying explanatory variable). Note that a positive coefficient indicates a higher probability of returning to employment, and hence a shorter duration of unemployment.

Whereas the unit of observation in the various duration models represented by equation (2) and displayed in Table 3 is the

8. If e is assumed to have the logistic distribution, then we have a logit model, which is prefered to a linear probability model because it yields unbiased, consistent, and efficient coefficient estimates. Exploratory work estimating equation (3) by logit has yielded results that are nearly identical--in terms of both statistical significance and quantitative response to variation in explanatory variables--to the linear probability estimates reported below.

claimant, the unit of observation in a hazard model is the claimant-week. The transformation of claimant records into claimant-week records is illustrated in Table 4. Panel A shows records for three claimants, who experienced 4, 38, and 0 weeks of insured unemployment. The first and third claimant became reemployed after receiving their last UI benefit payment, whereas the second remained jobless. Also, the first and third were ineligible for FCS, whereas the second was eligible.

The second panel of Table 4 shows the claimant-week records that are generated by the three claimant records in Panel A. The first claimant contributes a total of 6 observations to the claimant-week data set--one for the waiting week, one for each week in which UI benefits were received, and one more for the week following the spell of insured unemployment, since this worker became reemployed. The dependent variable in the hazard analysis, reemployment, is zero in all weeks except the last, in which reemployment occured. The second claimant contributes 39 observations to the claimant-week data set--one for the waiting week, and one for each week in which UI benefits were received. The reemployment variable is zero for all of these observations, and since this claimant did not find reemployment after exhausting his UI benefits, there is no fortieth observation following the spell of insured unemployment in which the reemployment variable equals one. Note that, when claimant records are transformed into claimant-week records, each claimant contributes exactly as much information as is known about him to the analysis of reemployment probability (Allison 1982).

Hazard models such as (3) start from the pioneering work of Cox (1972), and are often referred to as "semiparametric" because they implicitly make no assumption about the distribution of u in the duration equation (2). Several recent studies (for example, Moffitt 1985a, 1985b; Steinberg and Monforte 1987; Sheffrin and Dynarski 1987; Ham and Rea 1987) have estimated hazard models such as (3). Not all such studies have estimated the effects of increases in potential benefit duration on jobless duration: Moffitt's estimates suggest that a one-week addition to potential duration leads to a 0.15-week increase in joblessness.

The estimates displayed in Table 5 are from two possible specifications of equation (3). Specifically, we regress R_t on the same time-invariant individual characteristics as in the duration models, and add a set of dummy variables modeling the time since the initial claim (t_1, t_2, ..., t_N):

$$(4) \qquad R_t = h(t) = b_0 + b_1x_1 + \ldots + b_Kx_K + c_1t_1 + c_2t_2 + \ldots + c_Nt_N + u.$$

The time-since-initial-claim dummy variables (t_1, ..., t_n) can be thought of time-varying explanatory variables, since their values change as a claimant's time since filing for benefits lengthens. Note that the use of dummy variables allows estimation of a completely flexible relationship between the hazard rate and the time since filing the initial claim. No assumption is imposed on the shape of the hazard function.

A negative coefficient in the reemployment hazard model implies a lower probability of reemployment in any given week, and hence a longer expected unemployment duration. Thus, we expect variables with positive coefficients in the duration models to have negative coefficients in the Reemployment Hazard model. A comparison of the results in Table 5 with columns (3) and (4) in Table 3 shows that this is usually the case.

It is convenient to show graphically how the conditional probability of reemployment changes with the time since filing the initial claim, as in Figure 1. The reemployment hazards under misspecified censoring are shown as squares, whereas the hazards under corrected censoring are shown as plus signs.[9] Both hazard plots are surprisingly flat, and the differences between the two are not great except near the expiration points, which are 27 and 39 weeks after filing for benefits. That the hazards should differ at these points is important, however, because at these points erroneous assumptions are made under usual censoring conventions.

The results of main interest in Table 5 pertain to the effects of FSC. When censoring is misspecified--as in column (1) --it appears that the availability of FSC has no impact on the probability if returning to work. But under corrected censoring--in column (2)--the estimated impact of FSC availability on the probability of returning to work is large. The coefficient (-0.012) implies that jobless workers who were eligible to receive FSC were 1.2 percentage points less likely to return to work in any given week than were ineligible workers.

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9. The reemployment hazard rates are computed by substituting the characteristics of the average worker in each week into the hazard function, and solving for the dependent variable.

This 1.2 percentage point decrease in the reemployment probability needs to be viewed in context. The average weekly hazard--that is, the average probability of returning to work in any given week--was 0.035 or 3.5 percent for the average worker who was ineligible to receive FSC.[10] The availability of FSC lowered this probability to 0.023 (= 0.035 - 0.012), or 2.3 percent. This is a decrease in the probability of finding reemployment of about 33 percent, and suggests that FSC availability sharply reduced incentives to find reemployment.

An important limitation of the model specified as equation (3) is that it restricts the effect of FSC availability to be constant over the jobless spell. If FSC had an impact on the reemployment hazard rate that varied over the jobless spell, then this constant-proportion restriction would be undesirable.

It is straightforward to relax the constant-proportion restriction. Rather than enter eligibility for FSC as a single time-invariant dummy variable, FSC-eligibility can be interacted with the time-since-initial-claim variables to yield additional terms (t_1FSC, t_2FSC, ..., t_{N-1}FSC). These new interaction terms are then added to the model:

10. This so-called baseline hazard is computed by substituting the sample means into the estimated hazard function and solving for the dependent variable.

(5) $$Rt = h(t) = b_0 + b_1x_1 + \ldots + b_Kx_K + c_1t_1 + c_2t_2 + \ldots + c_Nt_N + d_1(t_1FSC) + d_2(t_2FSC) + \ldots + d_{N-1}(t_{N-1}FSC) + u.$$ [11]

The d_t coefficients will represent the change in the reemployment hazard in period t that is induced by the availability if FSC.

Table 6 and Figure 2 display the central results of estimating a model based on equation (5). Column (2) of Table 6, headed "FSC-Induced Change in Hazard," displays the coefficients of the interaction terms (d_1, d_2, ..., d_{N-1}), which again can be interpreted as the time-dependent change in reemployment hazard induced by FSC-eligibility. (The estimated coefficients come from a model identical to that displayed in column (2) of Table 5, with the addition of terms that interact FSC-eligibility with time-since-initial-claim.) The results suggest that FSC reduces the probability of reemployment late in the regular benefit spell (that is, in weeks 22, 23, and 26), and dramatically reduces the reemployment probability at the regular UI exhaustion point (week 27), and in the weeks during which FSC is received (weeks 28 through 35).

Figure 2 shows the reemployment hazard plots (that is, the conditional probabilities of reemployment over the spell of joblessness) for workers who are eligible for FSC and for those who are not.[12] The figure throws light on two points. First, the pattern the reemployment hazard for FSC-eligible workers

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11. Note that the Nth interaction term is omitted in order to avoid perfect multicolinearity.
12. As in Figure, these reemployment hazard rates are computed for the average worker who remains jobless in each week.

suggests that restricting the effect of FSC to be constant in all periods is highly misleading. Second, and more important, the figure illustrates the dramatic reduction in the reemployment hazard among FSC-eligible claimants at the time regular benefits expire (27 week after the initial claim). Indeed, the change in the reemployment probability induced by FSC in week 27 can only be described as astonishing: Claimants who exhausted their regular benefits and were ineligible for an additional 12 weeks of FSC had a probability of finding reemployment of 0.81; whereas claimants who exhausted their regular benefits but were eligible for FSC had a probability of reemployment of 0.13.

What are the implications for jobless duration of the findings displayed in Table 6 and Figure 2? It is a straightforward exercise to compute expected jobless duration from a complete hazard function (Ham and Rea 1987). Denote the reemployment hazard in week t as h(t). This is the probability of finding reemployment in week t, conditional on having experienced t-1 weeks of joblessness. It follows that the unconditional probability of experiencing t weeks of joblessness is:

$$p(t) = [1-h(0)][1-h(1)][1-h(2)] \ldots [1-h(t-1)][h(t)].$$

That is, p(t) is the probability of not finding a job in the first t-1 weeks (the product of the conditional probabilities of not finding a job in each of the first t-1 weeks) times the conditional probability of finding a job in week t [h(t)]. The expected jobless duration is then found as:

$$E(D) = \sum_{t=1}^{\infty} t[p(t)].$$

Carrying out this exercize suggests that the expected jobless duration of the sample mean worker would have been 40.3 weeks if he had been eligible for FSC, and 29.2 weeks if he had been ineligible for FSC. The difference, 11.1 weeks, is the implied effect of FSC on expected jobless duration. The result can be stated another way: Each week of FSC availability increased the expected duration of joblessness by over 0.9 week.

This result suggests that the effects of an additional potential week of UI benefits are significantly larger than most previous estimates have indicated. It is unclear whether the magnitude of the result obtained here should be attributed to one of more factors that are particular to these data--for example, to the fact that in these data FSC eligibility is uncorrelated with work history, or to the arguably improved specification of censoring that these data allow, or perhaps to peculiarities of the FSC program around the time it expired.

IV. Summary and Conclusions

This paper has explored the effects of extended unemployment benefits--specifically the Federal Supplemental Compensation program that existed from 1982 to 1984--on the length of unemployed workers' jobless spells. The data used come from the administrative records of the Illinois Department of Employment Security, and are especially well suited to such an inquiry for at least two reasons. First, because the data span the expiration of the FSC program at the end of 1984, they permit one to compare the jobless spells of workers who were actually eligible to FSC with the spells of workers who would have been eligible if they had become unemployed and filed for benefits a matter of weeks earlier. This differs from other available data sets, in which eligibility for extended benefits is not independent of individual characteristics such as base period earnings or weeks of employment prior to layoff. Hence, the Illinois data permit cleaner comparisions than data used in earlier studies, and we might expect more striking results.

Second, the Illinois data include earnings history data, which permit one to correctly classify spells of unemployment as complete or censored. Earlier researchers have freqently had to make what appear to be erroneous assumptions about whether a spell of insured unemployment represents a complete or censored spell of joblessness.

These data yield striking findings about the effects of FSC on the expected length of a jobless spell. A Weibull model of the duration of joblessness suggests that the availability of FSC increased the expected duration of a worker's jobless spell by nearly 5 weeks. The result can be stated another way: Each week of extended benefits increased the expected duration of joblessness by 0.4 weeks (= 4.8/12). This estimate is toward the upper end of the range of existing estimates of the effects of extended benefits on jobless duration.

A hazard model of the conditional probability of becoming reemployed yields a yet higher estimate of the influence of extended benefits on jobless duration. FSC eligibility dramatically lowered the probability of returning to employment around the time of benefit exhaustion and during actual receipt of FCS (Table 6 and Figure 3). It follows that FSC increased the expected length of unemployed workers' jobless spells by over 11 weeks. Stated another way, an additional week of potential benefits increased the expected duration of joblessness by nearly one week. The finding suggests that the effects of an additional potential week of UI benefits are significantly larger than most previous estimates have indicated.

Perhaps the most striking finding revealed in this work pertains to the effect that the availability of 12 weeks of extended benefits has on workers' probability of finding a job at the expiration of their 26 weeks of regular benefits. Claimants who exhausted their regular benefits and were <u>ineligible</u> for an additional 12 weeks of FSC had a probability of finding reemployment of 0.81; whereas claimants who exhausted their regular benefits but were <u>eligible</u> for FSC had a probability of reemployment of just 0.13. Hence, the results strongly suggest that workers tend to find jobs just as their UI benefits expire.

Two conclusions seem appropriate in view of the findings presented. First, it seems likely that previous research has had to make do with data that are rather poorly suited to estimating the effect of potential benefit duration on jobless duration. For example, in most data, variation in the potential duration of UI benefits is correlated with other factors such as earnings history. Also, many studies have been forced to make incorrect assumptions about which spells of unemployment were censored and which were complete. Either factor--correlation between potential benefit duration and worker characteristics or misspecified censoring--could lead to estimates of the effects of extended benefits on expected jobless duration that are downward biased. Accordingly, the results obtained here suggest that we might begin to revise upward our estimates of how much an additional week of benefit availability increases jobless spells.

The second conclusion is somewhat more negative: In order to obtain a convincing point estimate of the effects of extended benefits on jobless duration, it may well be necessary to conduct household surveys that follow-up known spells of insured unemployment so that completed spells of joblessness beyond the maximum duration of benefits can be observed. That is, administrative data have been pushed about as far as they can in this endeavor, and additional (very costly) data will be needed to advance our knowledge any further.

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Table 1

Classification of Workers by Weeks of
UI Benefits Claimed and
Subsequent Labor Force Status

Case	Number of Weeks of UI Benefits Claimed	Labor Force Status after Benefit Termination	Number of Observed Men in Illinois UI Data (proportion)
A	Maximum Potential	In Covered Employment	283 (0.13)
B	Maximum Potential	Out of Covered Employment	434 (0.20)
C	Fewer than Potential	In Covered Employment	1040 (0.48)
D	Fewer than Potential	Out of Covered Employment	405 (0.19)

Notes: Cases B and C are correctly characterized by usual censoring conventions; Cases A and D are misspecified by the usual conventions.

Table 2

Mean Insured Unemployment Durations for Men by Monetary and Temporal Eligibility for Federal Supplemental Compensation (FSC)

(Standard Errors in Parentheses)

Temporal Eligibility for FSC	Monetary Eligibility for FSC: Eligible	Monetary Eligibility for FSC: Ineligible
Eligible	21.387 (0.402) (N = 1131)	16.532 (1.144) (N = 79)
Ineligible	17.864 (0.318) (N = 866)	20.058 (0.957) (N = 86)
Difference	3.524 (0.513)	-3.527 (1.979)

<u>Notes:</u> In order to be monetarily eligible for FSC, a claimant needed have total base period earnings equal to at least 1.5 times earnings high-earnings quarter of the base period. To be Temporally Eligible FSC, a claimant needed to file an initial claim for UI benefits befor September 30, 1984. Insured unemployment duration refers to the tota number of weeks of benefits (both state regular and FSC) received in claimant's full benefit year.

Table 3

Alternative Models of Unemployment Duration

(Estimated Coefficients with Standard Errors in Parentheses)

Explanatory Variable	(1) Mean or Proportion (Standard Deviation)	(2) Linear OLS	(3) Weibull Models: Censoring Misspecified	(4) Weibull Models: Corrected Censoring
Constant	1.000 (0.000)	11.231 (1.857)	2.712 (0.169)	2.832 (0.166)
Age:				
20-24	0.196 (0.397)	--	--	--
25-34	0.449 (0.497)	0.958 (0.697)	0.120 (0.064)	0.053 (0.062)
35-44	0.234 (0.423)	2.273 (0.830)	0.237 (0.077)	0.196 (0.075)
45-54	0.122 (0.327)	2.584 (0.957)	0.343 (0.091)	0.299 (0.089)
Ethnicity:				
White	0.640 (0.480)	--	--	--
Black	0.252 (0.434)	4.530 (0.654)	0.477 (0.066)	0.372 (0.063)
Hispanic	0.086 (0.280)	2.382 (0.953)	0.237 (0.092)	0.196 (0.088)
Native American	0.009 (0.096)	5.060 (2.601)	0.413 (0.268)	0.209 (0.231)
Other Race	0.012 (0.111)	0.833 (2.238)	0.194 (0.222)	0.418 (0.249)
Number of Employers in Base Period	1.480 (0.754)	-1.114 (0.341)	-0.103 (0.031)	-0.071 (0.030)

Table 3
(continued)

	(1)	(2)	(3)	(4)
Base Period Earnings:				
< $6,000	0.211 (0.408)	3.622 (1.437)	0.450 (0.130)	0.246 (0.130)
$6,000-$18,000	0.499 (0.500)	4.740 (1.093)	0.515 (0.095)	0.318 (0.097)
$18,000-$30,000	0.222 (0.415)	2.822 (1.096)	0.321 (0.093)	0.200 (0.097)
> $30,000	0.068 (0.253)	--	--	--
Number of Referrals	0.205 (0.646)	0.224 (0.386)	0.004 (0.036)	-0.023 (0.034)
Dependents' Allowance Received	0.465 (0.499)	1.473 (0.529)	0.188 (0.050)	0.129 (0.048)
Weekly Benefit Amount:				
< $51	0.071 (0.257)	--	--	--
$51 - $120	0.313 (0.464)	1.127 (1.133)	0.110 (0.107)	0.087 (0.102)
> $120	0.616 (0.487)	1.125 (1.361)	0.108 (0.129)	0.077 (0.124)
Labor Market:				
Low UE, High Growth	0.101 (0.301)	-1.231 (0.867)	-0.096 (0.076)	-0.156 (0.074)

Table 3
(continued)

	(1)	(2)	(3)	(4)
Low UE, Stable Growth	0.080 (0.271)	-0.890 (0.968)	-0.120 (0.086)	0.040 (0.089)
Chicago (Average UE, Average Growth)	0.591 (0.492)	--	--	--
Average UE, High Growth	0.064 (0.245)	0.584 (1.054)	0.039 (0.097)	0.117 (0.097)
High UE, Some Growth	0.108 (0.311)	2.251 (0.855)	0.153 (0.082)	0.203 (0.080)
High UE, Little Growth	0.039 (0.193)	3.453 (1.320)	0.256 (0.124)	0.340 (0.129)
Eligibility for FSC:				
Monetarily and Temporally Eligible	0.523 (0.500)	3.293 (0.523)	-0.106 (0.051)	0.180 (0.048)
Monetarily Ineligible	0.076 (0.266)	0.070 (1.039)	-0.025 (0.105)	0.086 (0.098)
Monetarily Eligible, Temporally Ineligible	0.401 (0.490)	--	--	--
Weibull Shape Parameter	--	--	0.871 (0.020)	0.813 (0.020)
R-squared (adj)	--	0.078	--	--
F	--	9.339	--	--
Loglikelihood	--	--	-2,893	-2,683

Table 3
(continued)

	(1)	(2)	(3)	(4)
Number of Censored Observations	--	--	717	839
Total Sample Size	--	2,162	2,162	2,162

Notes: The dependent variable in column (2) is the number of weeks of UI compensated unemployment experienced during the benefit year. Mean of the dependent variable is 19.75 weeks. The Weibull models account for censoring of the dependent variable, imposing the assumption that the underlying distribution of unemployment spells is well-described by the Weibull distribution. See the text for a discussion.

Table 4

Transformation of Data on Claimants into

Data on Claimant-Weeks

Panel A: Claimant Records

Claimant	Weeks of Insured Unemployment	Reemployed	Weekly Benefit	Eligible For FSC
1	4	1	$149	0
2	38	0	$161	1
3	0	1	$128	0

Panel B: Claimant-Week Records

Claimant	Weeks of Since Initial Claim	Reemployed	Weekly Benefit	Eligible For FSC
1	0	0	$149	0
1	1	0	$149	0
1	2	0	$149	0
1	3	0	$149	0
1	4	0	$149	0
1	5	1	$149	0
2	0	0	$161	1
2	1	0	$161	1
2	2	0	$161	1
2	3	0	$161	1
.	.	.	.	.
.	.	.	.	.
.	.	.	.	.
2	37	0	$161	1
2	38	0	$161	1
3	0	0	$128	0
3	1	1	$128	0

Notes: The claimant is the unit of observation in the alternative models of unemployment duration in Table 3 (Panel A). The claimant-week is the unit of observation in the reemployment hazard models in Table 5 (Panel B).

This is new for MEA
Same as MSU version
minus columns 3+4.
(Table 5)

Table 5

Reemployment Hazard Models

(Estimated Coefficients with Standard Errors in I__

Explanatory Variable	(1) Censoring Misspecified	(2) Corrected Censoring
Constant	0.0889 (0.0116)	0.3178 (0.0099)
Age:		
20-24	--	--
25-34	-0.0041 (0.0023)	-0.0016 (0.0022)
35-44	-0.0079 (0.0027)	-0.0061 (0.0025)
45-54	-0.0115 (0.0031)	-0.0096 (0.0029)
Ethnicity:		
White	--	--
Black	-0.0155 (0.0021)	-0.0134 (0.0020)
Hispanic	-0.0086 (0.0031)	-0.0076 (0.0029)
Native American	-0.0130 (0.0080)	-0.0091 (0.0075)
Other Race	-0.0076 (0.0073)	-0.0130 (0.0069)
Number of Employers in Base Period	0.0038 (0.0012)	0.0027 (0.0011)

Table 5
(continued)

	(1)	(2)
Base Period Earnings:		
< $6,000	-0.0183 (0.0048)	-0.0103 (0.0045)
$6,000-$18,000	-0.0203 (0.0038)	-0.0135 (0.0035)
$18,000-$30,000	-0.0138 (0.0038)	-0.0099 (0.0036)
> $30,000	--	--
Number of Referrals	-0.0003 (0.0012)	0.0003 (0.0012)
Dependents' Allowance Received	-0.0066 (0.0017)	-0.0055 (0.0016)
Weekly Benefit Amount:		
< $51	--	--
$51 - $120	-0.0039 (0.0038)	-0.0030 (0.0036)
> $120	-0.0039 (0.0045)	-0.0024 (0.0042)
Labor Market:		
Low UE, High Growth	0.0043 (0.0030)	0.0070 (0.0028)

Table 5
(continued)

	(1)	(2)
Low UE, Stable Growth	0.0040 (0.0033)	-0.0011 (0.0031)
Chicago (Average UE, Average Growth)	--	--
Average UE, High Growth	-0.0015 (0.0035)	-0.0043 (0.0033)
High UE, Some Growth	-0.0058 (0.0028)	-0.0070 (0.0026)
High UE, Little Growth	-0.0096 (0.0041)	-0.0114 (0.0039)
Eligibility for FSC:		
Monetarily and Temporally Eligible	-0.0011 (0.0018)	-0.0121 (0.0017)
Monetarily Ineligible	0.0014 (0.0035)	-0.0027 (0.0033)
Monetarily Eligible, Temporally Ineligible	--	--
Time Since Initial Claim Variables	Included--see Fig. 1	Included--see Fig. 1
F	30.599	83.491

Table 5
(continued)

	(1)	(2)
Number of Censored Spells	717	839
Total Number of Spells	2,162	2,162

Notes: The dependent variable is a dummy variable equal to 1 if reemployment occured in week t, zero otherwise. The unit of observation is the claimant-week (rather than the individual claimant), and there are 47,014 claimant-week observations. See equation 3 in the text, and the accompanying discussion.

TABLE 6

Change in the Reemployment Hazard

Induced by Eligibility for FSC

by Time Since Initial Claim

(1) Time Since Initial Claim (Weeks)	(2) FSC-Induced Change in Hazard	(3) Standard Error
0	-0.0003	0.0067
1	0.0054	0.0067
2-3	0.0036	0.0049
4-5	0.0006	0.0051
6-7	0.0090	0.0052
8-9	-0.0021	0.0054
10	0.0064	0.0077
11	0.0026	0.0079
12-13	0.0001	0.0057
14-15	-0.0072	0.0058
16-17	-0.0061	0.0060
18-19	-0.0084	0.0062
20-21	-0.0099	0.0063
22-23	-0.0126*	0.0065
24-25	-0.0034	0.0067
26	-0.0317**	0.0098
27	-0.6831**	0.0133
28-29	-0.3522**	0.0270
30-31	-0.1064*	0.0555
32-33	-0.2412**	0.0782
34-35	-0.4973**	0.1105

<u>Notes</u>: Figures displayed in column (2) are coefficients from a hazard model similar to that in column (2) of Table 5, with the addition of variables interacting FSC eligibility with time since initial claim.

One asterisk (*) denotes rejection of the hypothesis that the FSC-induced change in reemployment hazard is zero at the 5-percent level of confidence; two asterisks (**) denotes rejection of the hypothesis that the FSC-induced change in reemployment hazard is zero at the 1-percent level of confidence.

Figure 1

Reemployment Hazard Rates for Men Under
Misspecified and Corrected Censoring

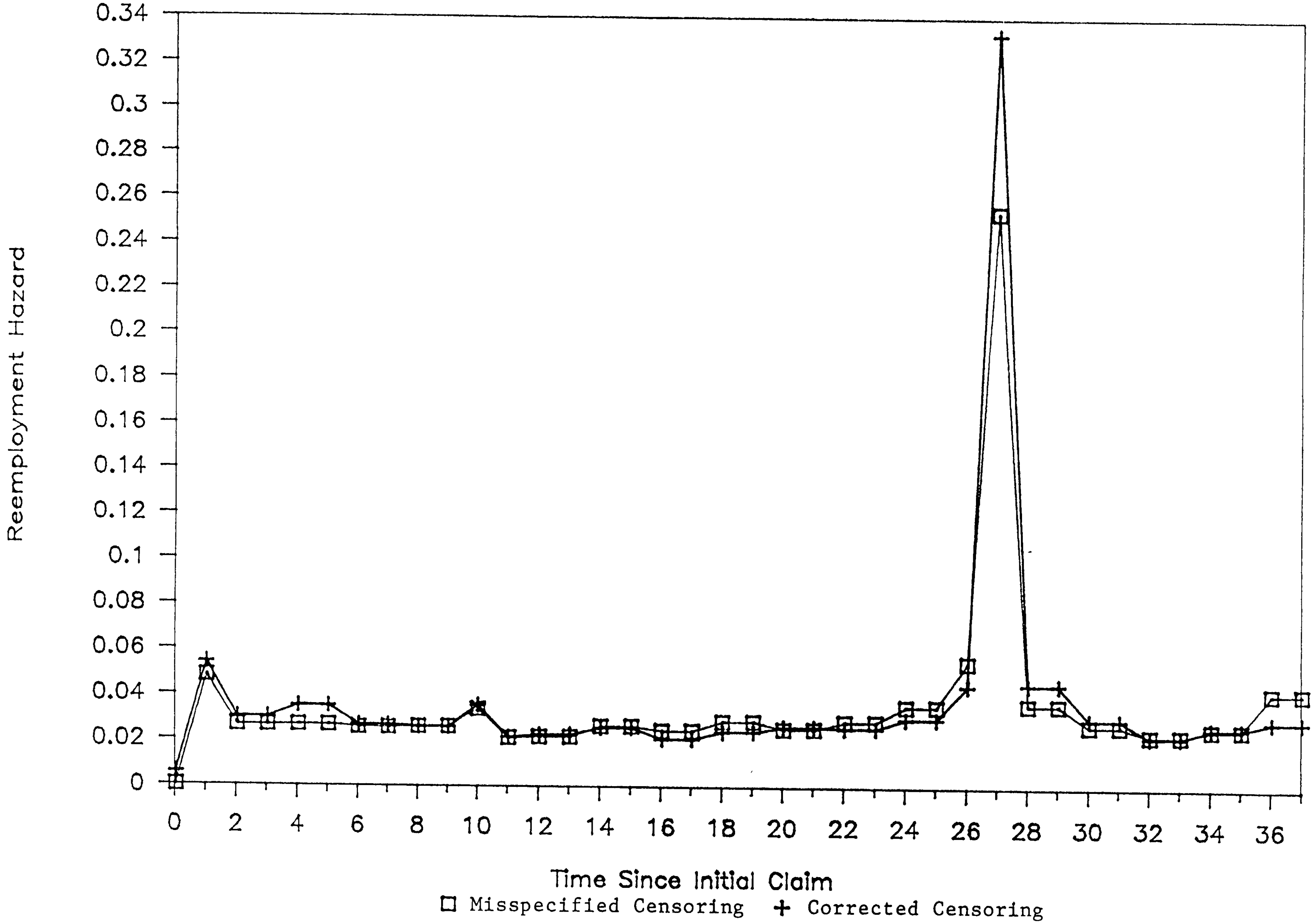

Figure 2

Reemployment Hazard Rates for Men who
were Eligible and Ineligible for FSC